What hides behind extended periods of youth unemployment in Bosnia and Herzegovina? Evidence from individual level data

(Preliminary Version)

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June 2008

Abstract

This paper provides the first empirical analysis on youth unemployment duration in Bosnia-Herzegovina. The study is based on micro data from the Household Survey Panel Series (2001-04). We formulate the problem within a duration model framework. Semi-parametric methods are used and compared to alternative approaches. The analyses are carried out separately for young men and women to take into account the traditional pattern of the domestic division of labour between genders. Our results indicate that the speed with which an unemployed young person finds employment is partly a function of his/her particular characteristics. We also find significant gender differences in factors affecting the prospects of access to employment. We further observe that for young men as well as young women there is strong evidence for non-monotonic duration dependence. These results turn out to remain robust to different specifications and to the introduction of unobserved heterogeneity.

JEL classification: J64, P20, J21.

Keywords: Unemployment duration, Transition countries, Labour market transition

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‡We would like to thank Jean-Claude Berthélémy, Jean-Pierre Laffargue and Furio Rosati for their helpful suggestions on the earlier version of this study. Any remaining errors are our responsibility.

1 Introduction

As a result of the break-up of Yugoslavia, armed conflicts throughout 1990s and ongoing economic and political reforms, the old economic structure and patterns of employment in Bosnia and Herzegovina[1] (BIH) have been challenged to their foundations. In this context, high rates of youth unemployment which have been accompanied by long spells of unemployment have emerged as one of the most serious problems for BIH.

An early discouraging experience in the labour market may have long-term consequences for young people's further integration into the labour market. Extended duration of unemployment may also be associated with social problems such as crime, violence, and drug abuse among others.[2] The length of time needed by unemployed young people to find employment opportunities appears to be a major concern to labour market policy analysts. The last decade has witnessed the publication of a large number of studies which have increased our knowledge of the determinants of individual unemployment duration in transition economies.[3] By contrast, far fewer papers study the determinants of outflows from unemployment into employment in BIH.

Young people face considerable difficulties in entering the labour market in BIH. This shows up in high youth unemployment rates, both in absolute terms and relative to the rate among older workers. In 2004, unemployment among young people was three times as high as in the adult population, indicating strong labour market disadvantages young people face compared to adults. Furthermore, youth unemployment is characterized by long-term unemployment. The proportion of unemployed young people who were looking for a job for more than a year amounted to over 70% in 2004.[4] Recent empirical studies present evidence of the magnitude and properties of the observed rates of labour market transitions (European Training Foundation, 2006; Fares and Tiongson, 2007 and Tiongson and Yemtsov, 2008). A comparison with other transition economies shows that, as far as the probability of moving from unemployment into employment is concerned, the labour market in BIH is characterised by low mobility for younger workers. A major finding of these studies is that the likelihood of being unemployed varies significantly across social space, with some social groups (for example, the least educated or young women) suffering more than others (the university educated or young men).

This paper provides the first empirical analysis on youth unemployment duration in BIH. The main goal of this empirical analysis is to examine various factors which cause

[1] The Dayton Peace Accords in 1995 established the two entities, namely the Republica Srpska (RS) and the Federation of Bosnia and Herzegovina (FBIH), that constitute together the sovereign state of BIH.

[2] For a study on youth vulnerability in South Eastern Europe, see La Cava *et al.* (2005).

[3] A comprehensive review of literature is provided in Svejnar (1999).

[4] Authors' computations from the Household Survey Panel Series Wave 4.

some young people (aged 15-29) to experience longer unemployment spells than others. We also want to shed some light on the dependence of conditional probabilities of leaving unemployment (hazard rates) on the duration of unemployment even when observable and unobservable differences across individuals are controlled for. Better understanding of unemployment duration differences between young men and women is a secondary goal. We formulate the problem within a duration model framework. The study is based on micro data from the Household Survey Panel Series (HSPS) which contain detailed information on households and individuals, including a retrospective unemployment history.

Non-parametric techniques are used to study the shape of the overall hazard function. These results also suggest the parametric specifications required to consider relevant individual characteristics. A particular accelerated failure time model is conducted in the first part of our econometric analysis. However, distributional assumptions used in the parametric analysis of duration data impose strong restrictions on the shape of the hazard function that frequently may be inappropriate. The study addresses these concerns by using a semi-parametric estimator in which the baseline hazard is modelled in a flexible way. The principal reason for using different specifications of the hazard function is to arrive at some conclusion regarding the relative merit of various methods of estimation.

Observed heterogeneity is incorporated with a set of variables on individual, household and local labour market characteristics. Estimates are carried out separately for young men and women to take into account the sexual division of labour in society and the different roles assigned to women and men. It is well known that omitted variables are likely to lead to a spurious duration dependence in the baseline hazard (Lancaster and Nickell, 1980). This unobserved heterogeneity will also potentially result in biases in the estimated effects of included covariates. Hence, we incorporate unobserved heterogeneity in the analysis by using a normal distribution.

The remainder of the paper is organized as follows. Section 2 describes the study's data base. Section 3 presents the econometric specification and Section 4 presents the main results. Section 5 concludes.

2 Data and variables

The data used in this analysis are drawn from the HSPS covering the period 2001-04. In the first wave (2001), a sample of 5,400 nationally representative households, approximately 9,400 individuals were interviewed. Half of the initial Wave 1 respondents were subsequently interviewed each year through 2004.[5] A unique identification number

[5] At each wave, sample persons include all selected Wave 1 residents who are still eligible, that is, have not emigrated or entered an institution, children as they become 15, and household members who leave to form or join new households. Where original sample members joined new households, all adult

for each person allows us to construct a four-period panel with 2,700 individuals aged 15 to 29. The questionnaire was designed to cover information on labour market status and also a range of topics including demographic characteristics, education, migration, and social assistance.

At each interview, detailed information on labour market outcomes was collected both for current status and retrospectively for the previous calendar year. The individual questionnaire used in the first wave contains a retrospective question, where unemployed respondents state how much time they spent in their current unemployment spell. At any subsequent wave, the data contain information on the labour market histories of the respondent the year prior to the interview. Respondents were asked to recall any labour market spells experienced since the last survey. For each spell, respondents provide information on its type (new job, self-employment, unemployment), its beginning and ending month. The duration of unemployment is measured as the length of time between the date of entry into unemployment and the date of the transition from unemployment into employment.

The most commonly used definition of unemployment is that based on ILO criteria.[6] While the general advantages of this definition are obvious, our procedure for the analysis of youth unemployment duration is to define unemployment in terms of respondents' self-report of their labour market position. The main reason for this choice is that the data set does not allow to compute consistent ILO unemployment indicators due to a routing error occurred in the 2003 questionnaire. Furthermore, the ILO approach raises a difficulty with respect to the issue of work motivation. In order to be considered as unemployed, people must have taken steps in the recent past to find employment. This tends to exclude unemployed young people who may become discouraged from active job search.

We analyse unemployment spells that started after June 1999,[7] up to November 2004. The ongoing unemployment spells at that time are censored. Only the ultimate spell is included for individuals experiencing multiple spells of unemployment to avoid the intra-person correlation. After these restrictions, the sample amounts to 780 unemployment spells among 410 young men and 370 young women.

members of the new household were added to the sample. As a result of these procedures, the attrition of the panel is low, around 4%. The sample is considered to reflect demographic changes and to remain representative of the population over time.

[6]Unemployment, adopting the ILO definition, refers to those who, during the reference week, were without work, were actively searching for work and available to start within two weeks.

[7]This restriction is due to end of the war in Kosovo and to the beginning of a relatively stable period in the region.

The empirical model controls for any observable variables likely to be related to a young person's of cost of unemployment, leisure-income preference, or distribution of employment opportunities. The variables include a number of characteristics such as age, ethnicity, entity of residence, migration status, and educational attainment. Heterogeneity in family composition is taken into account by introducing variables for marital status, presence of a child aged less than 15 in the household, presence of other unemployed person in the household and existence of a family-based establishment (proxied by a variable indicating that household head works on his own account either in agricultural or non-agricultural activities).

Some labour market related variables should also be introduced in the analysis: situation at the Employment Bureau (registered as unemployed or not registered) and whether or not the individual is a first-time job seeker. The monthly average unemployment rate of the municipality is controlled to proxy local labour demand conditions an unemployed worker faces. An indicator variable for receipt of social assistance benefits captures potential disincentive effects of alternative income sources on job search behaviour.[8] All the explanatory variables (except for local unemployment rate) are included in dummy variable form with one category of each excluded. Variable definitions and basic descriptive statistics are provided in Table 1.

In order to have a better picture of the data, non-parametric estimates (Kaplan-Meier) of the hazard function for young men and women are plotted in Figure 1 and Figure 2 respectively.[9] These estimates provide a preliminary idea on the shape of the hazard function without making any assumptions on its functional form. However, they do not allow to take account for either observable or unobservable heterogeneity. The completed spells of unemployment have a mean (median) duration of 21.6 (17.0) months with a standard deviation of 1.2 (0.5) months.

Empirical hazards reveal some indication of non-monotonicity and follow a flexible course. For both young men and women, hazard rates into employment are high in the early months of the sample. A significant spike (much more pronounced for young men than for young women) is noticed in the hazard function approximately at 15 months. Thereafter, the hazard rates fall. Hence, for both groups, the hazard appears inverse U-shaped with a downward sloping portion starting around 15 months. Given the large confidence intervals after 45th month, the second spike of the hazard for young men may be attributed to the relatively small sample size and may be considered spurious. It is also interesting, though not surprising, to note the striking difference in the hazard function

[8]The measure consists of the sum of social benefits received by the household such as, temporary and permanent allowances, old age pensions, disability pensions, veterans' pensions, war disability pensions and survivors' pensions.

[9]Vertical lines show 95% confidence intervals.

Table 1: Characteristics of the sample

Variable and Coding	Mean (Standard deviation)
Mean Duration of Job Search - Completed spells (DUR)	21.6 (1.2)
Median Duration of Job Search - Completed spells (DUR)	17.0 (0.5)
Age group	
Age 15-19 years (AGE 1) *reference group*	0.293
Age 20-24 years (AGE 2)	0.501
Age 25-29 years (AGE 3)	0.205
Ethnicity	
reference group: Serb origin (SERB)	0.389
Bosnian origin (BOS)	0.465
Croat origin (CRO)	0.144
Educational attainment	
reference group: No education (NODIPLO)	0.061
Primary education (PRIM)	0.237
Secondary education (SEC)	0.666
College education (COLL)	0.034
Migration status	
reference group: Never changed place of residence (NOMIGR)	0.542
Displaced because of armed conflicts (DISP)	0.330
Migrated for economic reasons (VOL)	0.162
Female (FEM)	0.474
Married (MAR)	0.248
Presence of children aged less than 15 in the household (CHILD)	0.391
Living in FBIH (FBIH)	0.538
First-time job seeker (FIRST)	0.670
Registered unemployment at the Employment Bureau (EMPBURO)	0.683
Presence of a family-based establishment (FAMEST)	0.090
Presence of another unemployed person in the household (HHUNEMP)	0.215
Is allowed for receipt of social assistance benefits (BENEF)	0.276
Monthly unemployment rate (%) at the municipality level (LOCUNEMP)	5.3 (2.0)
Sample size	**780**

Note: Standard deviation is provided only for continuous variables.

for men and women. The hazard function for young women lies entirely below that of young men.

Before turning to the econometric methodology, two major shortcomings of the data should be stressed. One limitation is the relatively small sample size. Due to small number of observations, transitions from unemployment into non-participation could not

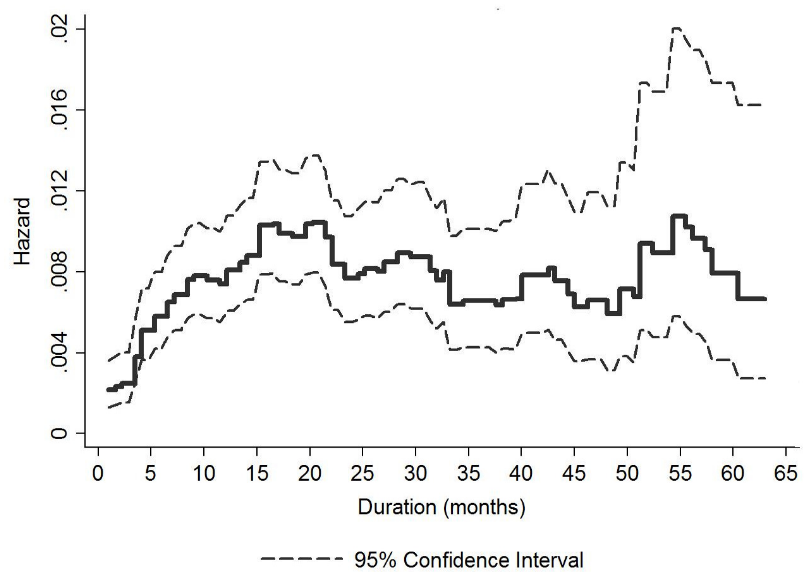

Figure 1: Empirical hazard estimates for young men

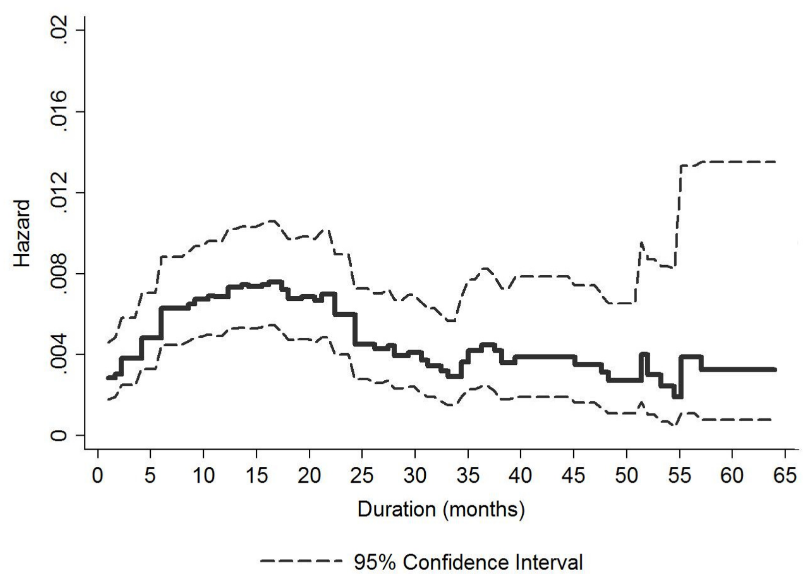

Figure 2: Empirical hazard estimates for young women

be modelled explicitly, but are treated right-censored. A second limitation is the lack of information on the amount of unemployment related benefits that individuals receive. In addition, information on the receipt of unemployment cash benefits is missing in 2002. Therefore, we choose to exclude this variable from the estimated model.

3 Empirical Implementation

The standard job search model provides a theoretical framework for modelling unemployment duration. Numerous models of unemployed workers' job search behaviour under imperfect information have been developed. For a review of job search models, see Devine and Kiefer (1991). The empirical analysis of unemployment duration is conducted by using a duration model.[10] Estimates are carried out separately for young men and women to take into account the differences in baseline hazards and other parameter values.

We specify an accelerated failure time (AFT) model where the effect of explanatory variables is to rescale time directly. We assume that the logarithm of unemployment duration T is linearly related to covariates x_i:

$$\ln T = \beta x_i + \varsigma T_0 \qquad (1)$$

where the random variable T_0 is scaled by a constant $\varsigma \equiv 1/\alpha$. In this study, we assume that the variable T_0 is specified as a gamma variable with parameter κ, which implies that the distribution of time to exit T is a generalized gamma[11] since the hazard may evolve over time in a non-monotonic way as suggested by non-parametric methods. The parameters β are estimated by a Maximum Likelihood technique. For additional discussion on the AFT model, see Kalbfleisch and Prentice (1980) and Lancaster (1990).

Fully parametric models of duration have two major drawbacks. First, the estimated effects of the covariates may be biased and misleading when a restrictive shape is imposed on the hazard rate of unemployed individuals. Second, the monthly data on unemployment duration available in the survey represent discrete observations of a continuous process. Hence, we estimate a semi-parametric version of our model along the lines of Prentice and Gloenker (1978), Meyer (1990) and Narendranathan and Stewart (1993).

[10]For a description of tools used in econometric duration models see Kiefer (1988), Lancaster (1990), Hosmer et Lemeshow (1999), and Horny (2006).

[11]The probability density function of a generalized gamma variate t is:

$$f(t,\kappa) = \frac{\alpha \eta^{\alpha\kappa} t^{\alpha\kappa-1} \exp\left\{-\left(\eta t\right)^{\alpha}\right\}}{\Gamma(\kappa)}$$

where $\Gamma(\kappa)$ is the gamma function. The generalized gamma nests the Weibull ($\kappa = 1$), the exponential ($\kappa = \alpha = 1$) and the log-normal distributions ($\kappa = 0$).

Formally, we parameterize the hazard for individual i at time t, using the proportional hazards (PH) form (Cox, 1972):

$$\lambda_i(t|x_i) = \lambda_0(t).\exp\left[x_i\left(t\right)\beta\right] \tag{2}$$

where λ_0 is referred to as the unknown baseline hazard function and $\exp\left[x_i\left(t\right)\beta\right]$ is the proportionality factor capturing the covariate effects.

The probability of a spell lasting until month $t+1$ given that it was still in progress in month t, is specified as a function of the hazard:

$$\Pr\left[T_i \geq t+1|T_i \geq t\right] = \exp\left[-\int_t^{t+1}\lambda_i\left(u\right).du\right] \tag{3}$$

Assuming that $x_i\left(t\right)$ is constant between t and $t+1$, Equation (3) can be rewritten as:

$$\Pr\left[T_i \geq t+1|T_i \geq t\right] = \exp[-\exp(x_i\beta + \gamma_t)] \tag{4}$$

where

$$\gamma_t = \ln\int_{t-1}^{t}\lambda_0(u).du \tag{5}$$

The logarithm of the integrated baseline hazards, γ_t, are treated as constants in each period and estimated along with the elements of β by Maximum Likelihood. Thus, the shape of the baseline hazard is estimated on a month by month basis without any further restrictions.[12]

In models considered so far, all differences between individuals are assumed to be explained by a vector of observed explanatory variables x_i. As pointed out earlier, duration models suffer from downward biased estimates and a spurious duration dependence when there is unobserved heterogeneity (Lancaster, 1990). We also consider a generalisation of equation 2 that adds an unobserved individual heterogeneity term:

$$\lambda_i(t|x_i) = \vartheta_i.\lambda_0(t).\exp(x_i\beta) = \lambda_0(t).\exp(x_i\beta + log(\vartheta_i)) \tag{6}$$

where ϑ_i follows a normal distribution with mean 0.

[12]Intuitively, the specification of the baseline hazard is a series of dummy variables which does not require an a priori distributional assumption. This is akin to the Cox partial likelihood estimation method (1972). However, instead of conditioning out the baseline hazard, we estimate it jointly with the coefficients of the covariates.

In this paper, the piecewise constant hazard (PCH) model is implemented by partitioning the time axis into 10 intervals. The model implies that the baseline hazard is assumed to be constant within each interval but different between intervals. The first 4 intervals are 3 months to capture the quickly changing hazard at the beginning of the spell. The next 5 intervals are 6 months and the last interval is a residual piece including 42 months and more.

4 Results

4.1 Estimated coefficients

The maximum likelihood estimates of the coefficients for young men and women are reported in Table 2 and Table 3 respectively. We can start by noting that the estimated coefficients turn out to remain robust to different specifications and to the introduction of unobserved heterogeneity. The AFT model and the PCH model generally yield similar results.[13] The inclusion of unobserved heterogeneity tends to increase slightly the absolute value of coefficients. For clarity, discussion of the determinants of unemployment duration concentrates on the results from the PCH model. The first column excludes ethnicity dummies because of their collinearity with the indicator variable for entity. They are included in the second column. Note also that for reasons of collinearity with interval-specific dummy variables (γ_i), we do not include the intercept term in the estimations of PCH model.

The coefficient estimates for the age dummy variables suggest that the younger the individual is, the better his/her prospects are of leaving unemployment. Young people aged 25 to 29 are less likely to take up employment as compared to their counterparts aged 15 to 19.

It seems that young women experience greater difficulties in finding employment when residing in FBIH than in RS. We add dummy variables to indicate whether or not ethnicity is a significant factor in explaining the flows to employment. Young women of Bosnian origin have significantly lower hazard rates to employment than their counterparts of Serb origin.[14] For the young male sample, we do not find any significant effect of entity or ethnicity.

[13]Given the difference in the specification of the dependent variable, the values of the coefficients in AFT model are not directly comparable with the PCH model coefficients. A positive coefficient in the PCH model implies that for changes in the value of a covariate the hazard rate increases while in AFT models a positive coefficient implies that the covariate is associated with longer duration times.

[14]The effect of ethnicity is interesting in comparison with studies in other transition countries. Ham *et al.* (1998) observe significantly longer unemployment durations for Romanies compared to other ethnicities in Czech and Slovak Republics. In Bulgaria, Romanies and Turkish ethnicities have a higher probability of being in long-term unemployment compared to Bulgarians (Kolev, 2003).

Table 2: Estimation results for young men

	AFT		**PCH**		**PCH unobserved heterogeneity**	
AGE2	0.352	0.376	-0.416	-0.424	-0.419	-0.427
	(0.276)	(0.297)	(0.273)	(0.272)	(0.275)	(0.275)
AGE3	0.640*	0.612*	-0.790*	-0.764†	-0.798*	-0.768†
	(0.313)	(0.306)	(0.393)	(0.394)	(0.399)	(0.399)
MAR	-0.488	-0.444	0.611†	0.579†	0.603†	0.561†
	(0.371)	(0.364)	(0.349)	(0.353)	(0.358)	(0.360)
CHILD	-0.006	-0.014	0.026	0.036	0.030	0.038
	(0.192)	(0.192)	(0.235)	(0.236)	(0.239)	(0.240)
FBIH	0.252		-0.321		-0.322	
	(0.201)		(0.230)		(0.234)	
BOS		0.192		-0.250		-0.239
		(0.197)		(0.236)		(0.240)
CRO		0.337		-0.444		-0.439
		(0.330)		(0.377)		(0.382)
PRIM	-0.921†	-0.761	1.124†	0.973	1.212†	1.067
	(0.567)	(0.649)	(0.775)	(0.784)	(0.781)	(0.789)
SEC	-1.578**	-1.409*	1.973**	1.820*	2.070**	1.923*
	(0.561)	(0.638)	(0.748)	(0.753)	(0.754)	(0.758)
COLL	-2.663**	-2.493**	3.316**	3.135**	3.397**	3.227**
	(0.548)	(0.586)	(0.988)	(1.001)	(1.004)	(1.015)
LOCUNEMP	0.570*	0.561*	-0.712*	-0.715*	-0.716*	-0.713*
	(0.267)	(0.265)	(0.316)	(0.315)	(0.322)	(0.320)
EMPBURO	0.395*	0.401*	-0.493*	-0.511*	-0.503*	-0.524*
	(0.188)	(0.180)	(0.221)	(0.222)	(0.225)	(0.225)
BENEF	-0.114	-0.111	0.140	0.139	0.146	0.146
	(0.192)	(0.188)	(0.237)	(0.238)	(0.241)	(0.243)
HHUNEMP	0.185	0.166	-0.243	-0.216	-0.247	-0.221
	(0.176)	(0.173)	(0.213)	(0.214)	(0.216)	(0.217)
FAMEST	-0.512†	-0.507†	0.618†	0.628†	0.640†	0.650†
	(0.324)	(0.318)	(0.363)	(0.367)	(0.370)	(0.372)
FIRST	0.147†	0.123†	-0.185†	-0.160†	-0.195†	-0.171†
	(0.092)	(0.076)	(0.114)	(0.101)	(0.121)	(0.106)
DISP	0.505†	0.507†	-0.647†	-0.649†	-0.675†	-0.687†
	(0.316)	(0.322)	(0.364)	(0.365)	(0.369)	(0.370)

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	AFT		**PCH**		**PCH unobserved heterogeneity**	
VOL	-0.117	-0.154	0.148	0.188	0.152	0.193
	(0.198)	(0.188)	(0.234)	(0.228)	(0.238)	(0.232)
Intercept	6.905**	6.737**				
	(1.088)	(1.127)				
γ_1			-8.343**	-8.236**	-8.480**	-8.358**
			(1.364)	(1.347)	(1.382)	(1.362)
γ_2			-8.701**	-8.592**	-8.834**	-8.710**
			(1.394)	(1.378)	(1.411)	(1.392)
γ_3			-8.408**	-8.295**	-8.536**	-8.409**
			(1.376)	(1.359)	(1.394)	(1.374)
γ_4			-7.660**	-7.547**	-7.785**	-7.657**
			(1.339)	(1.322)	(1.357)	(1.336)
γ_5			-7.354**	-7.238**	-7.467**	-7.335**
			(1.312)	(1.294)	(1.330)	(1.309)
γ_6			-7.410**	-7.295**	-7.514**	-7.383**
			(1.318)	(1.301)	(1.336)	(1.316)
γ_7			-8.064**	-7.948**	-8.161**	-8.029**
			(1.347)	(1.331)	(1.365)	(1.345)
γ_8			-7.583**	-7.460**	-7.677**	-7.538**
			(1.333)	(1.316)	(1.351)	(1.330)
γ_9			-7.983**	-7.860**	-8.072**	-7.934**
			(1.362)	(1.345)	(1.379)	(1.359)
γ_{10}			-7.578**	-7.455**	-7.661**	-7.524**
			(1.327)	(1.307)	(1.345)	(1.322)
ς	-0.187	-0.258				
	(0.426)	(0.560)				
κ	0.883	0.992				
	(0.641)	(0.855)				
N	407	407	12688	12688	12688	12688
Log-likelihood	-270.286	-270.324	-544.621	-545.062	-545.037	-544.180
Standard deviation of the unobserved heterogeneity variance					0.288	0.272
ρ (heterogeneity variance/1+heterogeneity variance)					0.048	0.043
Likelihood Ratio test of ρ=0						
χ^2 statistic					0.83	1.70

Note: See Table 1 for variable definitions
Standard errors are in parentheses. Significance levels : † : 10% $*$: 5% $**$: 1%

Table 3: Estimation results for young women

	AFT		PCH		PCH unobserved heterogeneity	
AGE2	0.387	0.397	-0.345	-0.355	-0.497	-0.503
	(0.321)	(0.319)	(0.339)	(0.341)	(0.465)	(0.464)
AGE3	0.798†	0.791†	-0.775†	-0.785†	-1.053†	-1.041†
	(0.430)	(0.424)	(0.459)	(0.459)	(0.640)	(0.643)
MAR	0.599†	0.571†	-0.577†	-0.561†	-0.815†	-0.772†
	(0.322)	(0.314)	(0.341)	(0.340)	(0.477)	(0.474)
CHILD	0.021	0.062	-0.149	-0.197	-0.034	-0.087
	(0.341)	(0.359)	(0.291)	(0.292)	(0.386)	(0.382)
FBIH	0.657*		-0.634*		-0.871*	
	(0.271)		(0.276)		(0.399)	
BOS		0.688*		-0.716*		-0.914*
		(0.284)		(0.296)		(0.422)
CRO		0.535		-0.493		-0.761
		(0.419)		(0.445)		(0.602)
PRIM	-1.687†	-1.600†	1.807†	1.850†	2.274†	2.176†
	(0.877)	(0.959)	(1.065)	(1.076)	(1.291)	(1.280)
SEC	-1.962*	-1.859*	2.119*	2.110*	2.697*	2.557*
	(0.842)	(0.910)	(1.048)	(1.061)	(1.293)	(1.292)
COLL	-3.115**	-2.938**	3.165**	3.075**	4.150**	3.899*
	(1.020)	(1.056)	(1.154)	(1.169)	(1.577)	(1.608)
LOCUNEMP	0.290	0.262	-0.257	-0.248	-0.416	-0.382
	(0.375)	(0.380)	(0.378)	(0.380)	(0.503)	(0.496)
EMPBURO	0.535†	0.552*	-0.622*	-0.630*	-0.731†	-0.748†
	(0.274)	(0.265)	(0.280)	(0.279)	(0.392)	(0.393)
BENEF	0.636†	0.663*	-0.580†	-0.620†	-0.852†	-0.879†
	(0.336)	(0.319)	(0.344)	(0.343)	(0.476)	(0.486)
HHUNEMP	0.733*	0.723*	-0.702*	-0.725*	-0.985*	-0.971*
	(0.288)	(0.281)	(0.289)	(0.289)	(0.416)	(0.424)
FAMEST	0.325	0.241	-0.243	-0.183	-0.455	-0.369
	(0.414)	(0.416)	(0.417)	(0.419)	(0.581)	(0.575)
FIRST	0.399	0.396	-0.360	-0.371	-0.537	-0.526
	(0.306)	(0.304)	(0.321)	(0.326)	(0.434)	(0.432)
VOL	0.013	0.047	-0.056	-0.085	-0.024	-0.061
	(0.376)	(0.375)	(0.377)	(0.379)	(0.498)	(0.491)

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... table 3 continued

	AFT		**PCH**		**PCH unobserved heterogeneity**	
DISP	-0.089	-0.100	0.098	0.123	0.114	0.123
	(0.306)	(0.311)	(0.316)	(0.316)	(0.410)	(0.405)
Intercept	5.736**	5.489**				
	(1.504)	(1.551)				
γ_1			-6.143**	-6.014**	-7.625**	-7.242**
			(1.624)	(1.634)	(2.237)	(2.265)
γ_2			-6.460**	-6.329**	-7.822**	-7.449**
			(1.649)	(1.659)	(2.204)	(2.217)
γ_3			-5.737**	-5.608**	-7.006**	-6.643**
			(1.616)	(1.626)	(2.143)	(2.149)
γ_4			-5.993**	-5.864**	-7.161**	-6.803**
			(1.636)	(1.647)	(2.130)	(2.128)
γ_5			-5.501**	-5.367**	-6.486**	-6.134**
			(1.596)	(1.607)	(2.061)	(2.048)
γ_6			-5.690**	-5.543**	-6.524**	-6.170**
			(1.607)	(1.618)	(2.042)	(2.023)
γ_7			-5.946**	-5.810**	-6.686**	-6.350**
			(1.627)	(1.638)	(2.044)	(2.023)
γ_8			-7.432**	-7.293**	-8.158**	-7.827**
			(1.871)	(1.878)	(2.241)	(2.220)
γ_9			-6.574**	-6.432**	-7.269**	-6.940**
			(1.734)	(1.742)	(2.124)	(2.101)
γ_{10}			-5.809**	-5.647**	-6.384**	-6.048**
			(1.631)	(1.638)	(2.028)	(2.003)
ς	0.365	0.327				
	(0.322)	(0.358)				
κ	0.225	0.292				
	(0.600)	(0.629)				
N	369	369	11091	11091	11091	11091
Log-likelihood	-196.37	-196.331	-352.525	-352.215	-351.975	-351.84
Standard deviation of the unobserved heterogeneity variance					1.443	1.384
ρ (heterogeneity variance/1+heterogeneity variance)					0.558	0.538
Likelihood Ratio test of ρ=0						
χ^2 statistic					1.10	0.75

Note: See Table 1 for variable definitions

Standard errors are in parentheses. Significance levels : $\dagger$: 10% $*$: 5% $**$: 1%

Our findings confirm the results of earlier studies with regard to educational attainment: education is crucial for the transitions from unemployment to employment for both sexes.[15] Workers with higher education experience the highest hazard to employment and those without any education (the reference group) have by far the lowest hazard to employment. We note that educational attainment is not only an indicator of productivity as desired by employers but also an indicator of labour market knowledge, greater job search efficiency, and the motivation to work. We further observe that the impact of educational attainment on unemployment duration increases with the level of education. Our results indicate also that the coefficient estimates are more drastic for women than for males indicating that the employment probabilities of women are more sensitive to education.[16]

Previous work experience reduces young men's job search time. Young people who have been previously employed have shorter unemployment spells than those with no previous experience. This can be explained by various factors. It is documented that the education system in BIH does not provide opportunities to combine initial education and work (Salines, 2007). Moreover, employers are often reluctant to hire young people for several reasons. They often want their employees to be immediately productive and they do not have enough financial and human resources to spend on the training of young people. Moreover, they find it risky to hire people without an employment record and work history (Kolev and Saget, 2005). For young women, we find that the probability of employment is not statistically affected by previous work experience.

Young people who are registered at the Employment Bureau tend to have longer unemployment spells. To explain this finding, one may surmise that they are more likely to benefit from passive and active labour market measures. However, employment agencies in BIH have a limited capacity to provide social welfare payments and services to the unemployed. According to World Bank (2002), new unemployment insurance systems introduced both in RS and FBIH are "modest, affordable systems of income support to the unemployed". The unemployment compensation schemes cover only a very small pro-

[15] See, for instance, Cazes and Scarpetta (1998) on Poland and Bulgaria, Voicu (2002) on Romania, and Kartseva (2002) on Russia. One of the important exceptions is Foley's study on Russia (1997) where better educated individuals do not find jobs more quickly than the less educated ones. This result is explained by the fact that education system inherited from the Soviet period is not suitable to the emerging market economy.

[16] Assuming some degree of professional skill specificity and variations in labour market conditions between different industries we would expect some variation in unemployment duration between young people specialised in different industries. However, in estimates not reported here, dummy variables controlling for professional skills turn out scarcely significant for both young men and women. Results are available upon request.

portion (less than 1%)[17] of young people registered as unemployed. In addition, given the significant number of young people who do not have any previous work experience, our sample consists mainly of new labour force entrants who are not eligible to receive unemployment-related benefits.[18] Hence, any incentive problems associated with the unemployment benefit system are unlikely to explain this hazard rate. Further research is required to ascertain more fundamental factors at work which could explain this pattern and the functioning of unemployment compensation system in BIH.[19]

How does migration status affect unemployment duration? Displaced young men are likely to have longer unemployment spells compared to those who never moved from their place of residence. There is no significant effect of voluntary (or economic) migration on unemployment durations. These findings suggest that the Bosnian labour market has been substantially affected by the recent conflicts and resulting movements of people. Similarly, Kondylis (2007) finds that displaced Bosnians are less likely to be in work. This result can be explained by various factors. First, unlike voluntary migrants, displaced people have had to leave their villages without being prepared for it, neither in material or psychological terms. All contact with their villages is cut off, since either the village is completely destroyed, or it is practically inaccessible. Consequently, displaced families are deprived of the initial material support from their villages of origin. Secondly, displaced people have had to migrate as whole families, whereas the former ones are part of chain-migration process, leaving the place of origin gradually as the pioneers settled down, established beneficial social networks, etc. Finally, absence of vocational skills and social networks that would help them to survive in a new environment and general marginalisation lead to difficulty in finding employment. For young women sample, migration variables do not seem to exert a significant effect.

[17] Authors' computations from the HSPS Wave 4.

[18] It is required for eligibility that applicants for unemployment benefits hold a considerable work record. Unemployment benefits recipients are required to hold a work record of at least 8 months within the last 12 months.

[19] According to theoretical and empirical analyses in the job search literature, the provision of unemployment benefits may have a negative effect on the labour market. Unemployment benefits tend to increase unemployment durations by raising individuals' reservation wages and decreasing the amount of time or effort to searching for a job (Atkinson and Micklewright, 1991). Studies on how unemployment benefits affect unemployment duration in other transition countries have not come up with consistent findings. Boeri (1997) and Boeri and Terrell (2002) argue that the initial generosity of unemployment compensation schemes may be an important factor in explaining the increase in registered unemployment at the beginning of the transition process. Jones and Kato (1997) show that in Bulgaria the receipt of unemployment benefits has a significant negative effect on women's reemployment probabilities. Ham *et al.* (1998) find that the unemployment compensation system in Slovak and Czech Republics has a moderately negative effect on the exit rate from unemployment. There are also a number of studies, including Lubyova and Van Ours on Slovak Republic (1997) and Puhani on Poland (2000), that find no evidence of such disincentive effects.

Regarding the influence of household-related variables, they have different effects on exit from unemployment among men and women. Being married decreases the probability of finding a job for young women while it speeds up the process of employment for young men. The prevailing pattern of the sexual division of labour within households may be an important factor in explaining this result. Having a spouse increases the propensity to find a job only among unemployed men because married men are considered as the main breadwinner in the household whereas women are largely expected to be full-time housewives when married. Generally, married men are under greater financial pressures to find a job. The burden of domestic responsibilities reduces the amount of time that women can devote to their job search, and their willingness to accept any kind of job. Accordingly, we may consider that, in BIH, men's and women's behaviour in the process of finding a job is largely dependent on their family status. Interestingly, the presence of children (under the age of 15) does not have affect unemployment duration significantly for men or for women, even though it has the expected negative effect on young women's probability to leave unemployment. As a robustness check, an interaction term is added to see if the presence of children matters more for young married women.[20] We found no evidence of any important interaction effects. These findings on the presence of children in the household are in line with those obtained by Foley (1997) for Russia and Lubyova and Van Ours (1998) for Slovak Republic.

The other household characteristics that affect young men's likelihood of leaving unemployment include the existence of household-based establishments. We find that there exists a significantly positive influence of a family enterprise or farm in job-finding probability of young men. We find no similar effect for young women. The presence of another unemployed person in the household results in lower hazard rates for young women. This result can be explained by the effect of this variable on the flow of information entering in the household and the knowledge that an individual may have of the job opportunities available in the labour market. Both results suggest that family connections and parental influence are among the important factors. This seems relevant in our context since job search in BIH is highly dominated by the use of social networks and family.[21]

For young women, if the household receives social benefits, it results in a significantly lower probability of leaving unemployment and hence longer unemployment duration.[22] However, our results do not lend strong support to the hypothesis that social benefits act

[20] Detailed results are not presented here due to space limitations and are available upon request from authors.

[21] For young people, the most frequently used method of job search is checking with friends and relatives. This is also the most productive method, in terms of offers and acceptances generated. In fact, this informal method accounts for almost 50% of jobs obtained by young people (authors' computations from the HSPS-Wave 4).

[22] In estimates not reported here we further observe that married women who receive social benefits are less likely to find an employment.

as a disincentive to exit unemployment for young men. We conclude that women are more sensitive to the incentive effects of social benefit systems. Alternative sources of income may push up young women's reservation wages and lower their job search intensity. In societies with prevailing traditional gender roles, women may prefer, or feel a greater social pressure, to emphasize their domestic roles. Reduced financial constraints may make it easier for them to give a lower salience to employment (Alm and Gallie, 2000).[23]

Finally, the general condition of the local labour market proxied by the monthly municipality unemployment rate has the expected signs. Living in municipalities with higher unemployment rates tends to reduce the probability of finding employment for young men. A deterioration of local labour market conditions reduces the probability of finding an acceptable job offer for any given amount of search, and thus lengthens the duration of unemployment. However, we do not find evidence that similar labour demand shocks have an impact on the employment probabilities of women.

4.2 Duration dependence

The AFT model with a generalized gamma distribution incorporates the most widely used distributions as special cases (exponential, weibull, log-normal). The estimated values for parameters κ and ς are significantly different from unity. Hence, we reject all the restrictions corresponding to the special cases mentioned earlier. To reduce potential bias in the estimation of the duration dependence pattern, estimates of the baseline hazards concentrate on the PCH model (Table 2 column 3 for young men and Table 3 column 3 for young women). The baseline hazards emphasize the gender differences in exit rates from unemployment into employment. We observe that young women have lower hazards than young men.

The results confirm the picture given by Figures 1 and 2. The duration dependence pattern for unemployment duration in BIH is non-monotonous. We see an initial rise amongst young people in the first 12 months. Thereafter the hazard rate falls following a downward course, which implies a negative duration dependence after the spike around the 12th month. A possible explanation of this negative duration dependence in the theoretical framework is that the arrival rate of job offers declines so much that it more than offsets any possible decline in the reservation wage. The net results will then be that the hazard function decreases with the duration of unemployment, which means that the longer an individual is unemployed the lower is the conditional probability of leaving unemployment.

[23]This large disincentive effect of alternative income resources has been suggested by Kupets (2005) for Ukraine, who pool men and women rather than examining unemployment durations by gender and show that income from pension has a strong disincentive effect on the probability of finding an employment.

This empirical result indicates a scar or stigma effect of unemployment in the sense that individuals with long unemployment durations will experience more difficulties in finding a job than otherwise identical individuals with shorter unemployment durations. There are several factors that could lead to the generation of this stigma effect. Extended unemployment can lead to human capital and working skills deterioration and loss of social networks. Unemployed youth may become discouraged by their lack of success in finding work and begin to search for work less intensively. Furthermore, prospective employers may take a record of extended unemployment as an indicator of low productivity.

While this finding may be due to true duration dependence, it may also be explained by a weeding-out effect once one acknowledges the presence of unobserved heterogeneity across individuals. If some individuals, because of unobserved factors (motivation, ability and so forth), have lower employment probabilities than other seemingly identical individuals, they will tend to stay unemployed longer. Then, even if individuals' employment hazards are constant over time, the data will display spurious duration dependence. In no case of this study is the hypothesis of no unobserved heterogeneity rejected (the likelihood ratio tests do not result in rejections), which suggests that heterogeneity not controlled for is being captured by flexible baseline hazard function. A possible explanation is that the identification of the effects of unobserved individual heterogeneity from duration dependence typically requires functional form assumptions that could not be tested in practice. Hence, in addition to the results reported here, an alternative functional form, namely a gamma distribution following Meyer (1990), has been tried. The results were generally unchanged.[24]

It may be interesting to compare our results on the presence of duration dependence and unobserved heterogeneity to those in empirical studies that use data from other transition economies. The results are mixed, since they depend heavily on the parametric specification for the duration dependence pattern and the unobserved heterogeneity distribution. Our results on duration dependence are fairly similar to those obtained in Foley (1997) and Grogan and Van den Berg (2001) for Russia. They use a flexible piecewise constant specification for the duration dependence incorporating an unobserved heterogeneity term. They find evidence for strong non-monotonous duration dependence, namely first increasing and then decreasing. The evidence for the presence of unobserved heterogeneity is rather weak.

[24] A non-parametric specification for the unobserved heterogeneity term is suggested by Heckman and Singer (1984) to correct the bias. However, Baker and Melino (2000) reveal a trade-off between the flexibility of the baseline hazard and the number of the mass-points used in the non-parametric unobserved heterogeneity distribution. They indicate that the non-parametric specification of both the distribution of duration and that of unobserved heterogeneity can lead to large and systematic biases.

5 Conclusion

This paper presents the results of an econometric analysis of the conditional probability of leaving unemployment, and hence of the determinants of individual unemployment durations, for young men and women living in BIH. Our statistical analysis demonstrates that the speed with which an unemployed young person finds employment is partly a function of his/her particular characteristics. We further observe significant gender differences in factors affecting the prospects of access to employment. The social assistance policy framework, and the societal gender role attitudes seem extremely important in influencing young people's duration of unemployment. The analysis also indicates that, after correcting for the observable and unobservable characteristics, the speed and likelihood of finding employment are significantly related to the total time spent out of work. An important finding is that the conditional probability of an individual finding a job declines steadily after 12 months of a spell, falling to a rather low level. Neither the sign nor the significance of the parameters is sensitive to different specifications and to the introduction of the unobserved heterogeneity.

All of these results point out to an agenda for future research. The availability of a longer panel data would improve the precision of estimates. Reliable empirical evidence on how unemployment compensation and social assistance schemes affect the job search behaviour of unemployed workers is crucial for getting a better grasp of the determinants of youth unemployment duration in BIH.

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